

GOVERNMENT PAYMENTS AND FARM BUSINESS SURVIVAL

NIGEL KEY AND MICHAEL J. ROBERTS

Using farm-level panel data from recent U.S. Agricultural Censuses, this study examines how direct government payments influence the survival of farm businesses, paying particular attention to the differential effect of payments across farm-size categories. A Cox proportional hazards model is used to estimate the effect of government payments on the instantaneous probability of a farm business failure, controlling for farm and operator characteristics. Results indicate that an increase in government payments has a small but statistically significant negative effect on the rate of business failure, and the magnitude of this effect increases with farm size.

Key words: agricultural payments, duration analysis, farm business survival, hazard model.

Economists and policy makers have long been interested in how government payments influence the growth and survival of farm businesses (e.g., Leathers, 1992; Tweeten, 1993; Atwood, Watts, and Baquet, 1996; Huffman and Evenson, 2001). With government payments to farmers exceeding \$20 billion in 1999, 2000, 2001, and 2003, farm payments have received greater public scrutiny, with some maintaining that payments unfairly advantage large operations (e.g., Williams-Derry and Cook, 2000; Becker, 2001). These concerns spurred congressional efforts to tighten payment caps on large-scale producers during the 2002 Farm Act debate (e.g., Nelson, 2002). The effect of payments on farm survival continues to be an important issue in on-going international trade negotiations, where agricultural support programs are a major source of contention.

This study uses a unique limited-access farm-level panel data set created from the 1987, 1992, and 1997 Censuses of Agriculture to derive the first estimates of the effect of government payments on the survival of individual U.S. farm businesses. Specifically, we use a Cox proportional hazards model to estimate the effect of government payments on the instantaneous probability (hazard rate) of a farm business failure. We consider the sur-

vival of individual businesses, controlling for the size, location, and organizational structure of the operation, and the age, race, sex, and career specialization of the operator. Separate estimates of the effect of payments on business survival are obtained for four farm-size categories.

This study exploits an exogenous source of variation in government payments—differences in payments that result from differences in “base acreage” in otherwise similar farms. Farmers that operate the same amount of land, located in the same county, producing the same crop may receive different levels of government payments if they have different amounts of land enrolled as “base acres”—land enrolled in a particular commodity program based on past plantings. Prior to 1996, acreage reduction provisions and restrictions on what could be planted on base acreage discouraged some farmers from fully participating in government programs—between 15% and 40% of eligible cropland was not enrolled in a Federal program (USDA-ERS). Due to historical variation in participation, similar farms had different base acres and received different amounts of government payments.

We find that government payments have a small but statistically significant negative effect on the instantaneous farm business failure rate. We also find that government payments reduce the failure rate proportionally more for larger farms. These results suggest that past agricultural support payments have contributed disproportionately to the survival of large operations.

Nigel Key and Michael J. Roberts are agricultural economists at the Economic Research Service.

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Related Literature

There is a substantial theoretical and empirical literature relating to firm size and firm survival. Jovanovic (1982), Ericson and Pakes (1992), and Pakes and Ericson (1998) present models in which firms (or entrepreneurs) are uncertain about their own efficiencies at startup. In these models, firms gradually learn about their abilities over time. The longer a firm operates in the market, the more information is gained. Firms that revise their perceptions of their ability upward over time tend to expand, while those revising downward tend to contract or exit. Consequently, the longer a firm has existed, the bigger it will be and the less likely it will be to fail. Empirical studies generally confirm these theoretical predictions (Dunne, Roberts, and Samuelson, 1988; Baldwin and Gorecki, 1991; Audretsch, 1991; Audretsch and Mahmood, 1995; among others).

For small businesses, the personal characteristics of the owner, such as educational attainment, can be important for small business survival (Bates, 1990; Taylor, 1999). The operator's age may be another important determinant of firm size and survival. Age may be correlated to knowledge about the firm's competitive abilities—with older owners able to acquire more information (Jovanovic, 1982). Alternatively, the operator's age may be related to financial liquidity. In the presence of liquidity constraints, it may take many years for business owners to accumulate sufficient net worth to obtain a certain scale of production (Evans and Jovanovic, 1989; Holtz-Eakin, Joulfaian, and Rosen, 1994).

Government payments could influence farm business survival through a variety of mechanisms. Farms receiving high payments per acre could bid up prices of fixed resources—especially land—causing low payment-per-acre farms to shrink or exit. Payments could also influence farm survival through capital market mechanisms. Government payments effectively raise a farm's net worth. This could make it less costly for the farm to obtain financing when liquidity constraints cause a farm's cost of capital to depend on its net worth (Hubbard, 1998). If large farms are liquidity constrained and small farms are not, then an increase in payments per acre can cause large farms to expand and increase in number, which bids up land prices causing small farms to shrink and decline in number (Key and Roberts, 2005). Higher payments may also

make agriculture more profitable relative to alternative occupations, which could reduce the incentive to exit farming.

Although a limited number of econometric studies have attempted to explain changes in the size and survival of individual farms based on characteristics of the farm operator or farm (Sumner and Leiby, 1987; Hallam, 1993; Zepeda, 1995; Weiss, 1999; Kimhi and Bollman, 1999), none have considered the role of government payments. A few studies have examined the relationship over time between government payments and *aggregate* measures of farm structure, including the national agricultural bankruptcy rate (Shepard and Collins, 1982), the total number of farms (Tweeten, 1993), and average farm size (Huffman and Evenson, 2001). More recently, Dixon et al. (2004) used state-level panel data to examine how government payments and other factors influence Chapter 12 bankruptcy filing rates. To our knowledge, this is the first study to examine the effect of government payments on the survival of individual farms.

Data

The data used in this study are from the Census of Agriculture files maintained by the National Agricultural Statistics Service.¹ The Census is conducted every four or five years and includes all U.S. farms. While offering a very large population, the Census does lack some information that would be useful for this study, such as operator educational attainment, and detailed financial information, such as return on assets or debt-coverage ratio. Since we are interested in the effect of government payments on farm survival, and to reduce sample heterogeneity, we restrict our analysis to "program crop farms"—those operations with Standard Industrial Classification (SIC) codes indicating specialization in wheat, corn, soybean, rice, cotton, or "cash grains."² Farms with these six SIC codes receive the largest shares of government farm payments.

Data from the 1978–97 Censuses can be used to examine farm survival rates by SIC

¹ More information about the Census of Agriculture can be found at: <http://www.nass.usda.gov/census/>.

² SIC codes for wheat, corn, soybean, or rice are assigned if any one of these crops account for at least 50% of sales. An operation is classified as a "cash grain" farm if a combination of these crops, or another cash grain not elsewhere classified totals at least 50% of sales.

Table 1. New Program Crop Farm (1982) Survival Rates over Time by Farm Type

Farm Category	1982	1987	1992	1997
All program crop farms				
Number surviving	140,876	70,478	45,122	31,630
Survival rate		(50.0)	(32.0)	(22.5)
Wheat (SIC = 111)				
Number surviving	20,592	10,534	6,678	4,697
Survival rate		(51.2)	(32.4)	(22.8)
Rice (SIC = 112)				
Number surviving	1,750	864	525	330
Survival rate		(49.4)	(30.0)	(18.9)
Corn (SIC = 115)				
Number surviving	46,150	23,091	14,876	10,363
Survival rate		(50.0)	(32.2)	(22.5)
Soybean (SIC = 116)				
Number surviving	34,875	15,398	9,311	6,392
Survival rate		(44.2)	(26.7)	(18.3)
Cash grain (SIC = 119)				
Number surviving	32,643	18,330	12,396	8,927
Survival rate		(56.2)	(38.0)	(27.3)
Cotton (SIC = 131)				
Number surviving	4,866	2,261	1,336	921
Survival rate		(46.5)	(27.5)	(18.9)

Note: The survival rate (in parentheses) is defined as the number farms surviving in a given period, as a percentage of the total number of new program crop farms established in 1982.

code.³ Table 1 presents the survival rates by SIC code for program crop farms that were *first observed* in the 1982 Census (these farms initiated production sometime between 1978 and 1982, as 1978 was the year of the previous Census). About 50% of new farms failed within the first five years. After ten years, about 32% of farms remained in business, and after fifteen years 22.5% remained in business. These survival rates are comparable to what has been reported for non-agricultural firms (e.g., Audretsch, 1991; Mata, Portugal, and Guimaraes, 1995; Disney, Haskel, and Heden, 2003). Consistent with past studies, the probability of survival generally increases with the age of the firm (Evans, 1987a,b; Audretsch, 1991). Cotton and rice farms had somewhat lower than average and cash grain farms somewhat higher than average fifteen-year survival rates.

Because of the way information in the Census of Agriculture is collected, this study focuses on the duration of a farm business continuously operated by the same individual. The Census collects information as to when the cur-

rent operator began to operate the farm, but not about how long the farm has been operating. Hence, there is no way to estimate the life of a farm business unless the same operator manages it.⁴ Consequently, we define a surviving farm as one remaining in business and having the same operator; farms remaining in business with a different operator were removed from the sample.⁵

This study examines the survival of program crop farms operating in 1987—the first year the Census of Agriculture began collecting information on government payments. Our sample includes the 200,187 program crop farms that had at least 10 acres of land and \$10,000 in sales in 1987 and for which information on

⁴ The Census tracks operations over time using a Census File Number (CFN). The Census defines a farm as out of business if there is no response to the Census questionnaire or the questionnaire is returned stating that the farm is no longer operating. However, if a farm changes operators through a business transaction or inheritance, the CFN may change even though the business is still operating. Hence, it is not possible to estimate the duration of a farm business based on how long the CFN appears in the Census.

⁵ We consider a farm to have the same operator if the age of the operator differs by five years between consecutive Censuses. About 8% of continuing farms had operators whose age differed by more or less than five years, and were therefore eliminated from the sample.

³ Individual Census data are not available prior to 1978.

Table 2. The Average Farm Business Lifespan by Sales and Government Payments as a Share of Sales

Sales Quartiles	Government Payments as a Share of Sales (θ) – Quartiles				
	Q1 ($0 < \theta < 0.12$)	Q2 ($0.12 \leq \theta < 0.22$)	Q3 ($0.20 \leq \theta < 0.36$)	Q4 ($\theta \geq 0.36$)	Q4 – Q1
Q1 (\$10,000 ≤ sales < \$23,991)					
Years	25.37	25.22	26.24	27.87	2.50***
(SE)	(0.119)	(0.163)	(0.157)	(0.124)	(0.172)
Obs.	17,031	8,615	9,253	15,145	
Q2 (\$23,991 ≤ sales < \$50,600)					
Years	25.94	26.60	28.00	28.48	2.54***
(SE)	(0.137)	(0.136)	(0.129)	(0.120)	(0.182)
Obs.	12,130	11,153	12,320	14,441	
Q3 (\$50,600 ≤ sales < \$104,390)					
Years	26.04	27.45	28.66	28.28	2.24***
(SE)	(0.138)	(0.137)	(0.114)	(0.117)	(0.181)
Obs.	9,952	13,343	13,642	13,114	
Q4 (sales ≥ \$104,390)					
Years	26.16	27.80	28.03	28.29	2.13***
(SE)	(0.156)	(0.077)	(0.083)	(0.118)	(0.195)
Obs.	6,141	24,039	21,994	11,696	

Note: Three asterisks (***) indicate that the null hypothesis of equal mean lifespan for the first and fourth payments-as-a-share-of-sales quartiles is rejected at the 0.001 significance level.

all variables was available.⁶ The Census allows us to identify whether a farm business ceased operating between 1987 and 1992, or between 1992 and 1997, or whether it was still operating in 1997. In addition, the Census records the year in which the current operator began managing the operation. Therefore, the observed lifespan of the farm business is defined as 1987 minus the year the operator initiated farming on the operation plus 2.5, 7.5, or 10, depending on whether the operation failed by 1992, failed by 1997, or remained in business in 1997, respectively. If the operation remained in business in 1997, the lifespan is right censored.

Because of the way we define the age of the business, all lifespans are left truncated. We do not begin to observe businesses until 1987—a known time after they began operating, and the risk set does not include businesses that failed prior to 1987. For example, of all businesses initiated in 1980, we only observe those businesses in 1987 that survived at least seven years. We do not observe farms that failed before 1987. Hence, for businesses that began in 1980 the lifespan is left truncated at seven years. The observed lifespan is there-

fore conditional on the period of truncation being exceeded.⁷

Methods and Results

To examine the relationship between government payments and farm business survival we first compare the average observed lifespan for farm businesses of different sizes and different shares of government payments in total sales. Table 2 shows that, with few exceptions, within each sales quartile, a larger share of government payments in sales corresponds to a longer average lifespan.⁸ For example, for farms with more than \$104,390 in sales, those farms where payments comprise less than 12% of sales have an average lifespan of 26.16 years compared to 28.29 years for those farms where payments comprise more than 36% of sales. The last

⁶ Deleting farms with less than \$10,000 in sales (which represent about a fifth of the observations) focuses the analysis on farm households where farm business income is a larger share of total income and where government payments are likely to play an important role in the decision to continue farming.

⁷ Left truncation is accounted for in the estimated likelihood function associated with the Cox proportional hazard model and the product-limit survival function estimates discussed in the next section (see SAS 9.1 PHREG Procedure, p. 2998, for details).

⁸ The average observed lifespans reported in table 2 do not account for left truncation or right censoring, meaning these averages are not unbiased estimates of the true lifespans. The standard error (SE) for the difference between any two average lifespans can be calculated by taking the square root of the sum of the individual squared standard errors ($\sqrt{SE_1^2 + SE_2^2}$). For example, the difference in average lifespans for the first sales quartile and second and third payments quartiles is $26.24 - 25.22 = 1.02$, and the SE for this difference is $\sqrt{1.57^2 + 1.62^2} = 0.226$, which gives a *t*-statistic of 4.5.

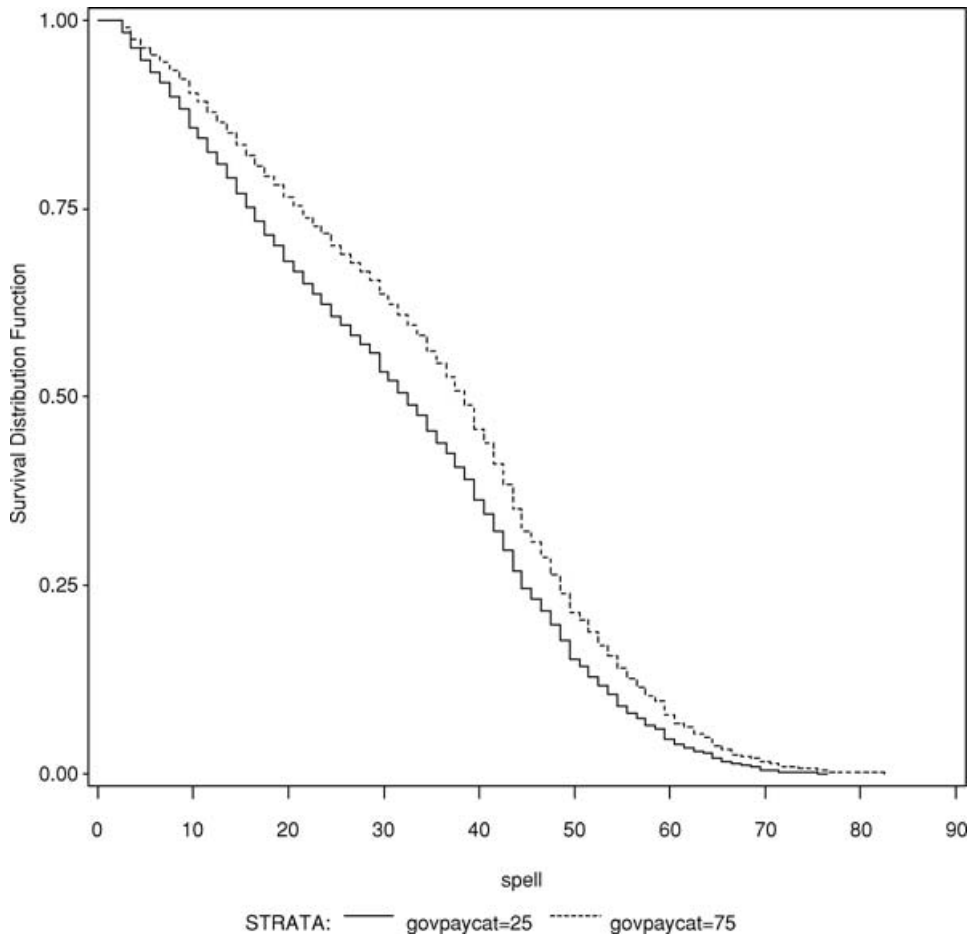


Figure 1. Kaplan-Meier estimated survival functions for farms in the upper and lower government-payments-as-a-share-of-sales quartiles

column in the table shows the difference in the average lifespan for farms in the first and fourth payments-as-a-share-of-sales quartiles. A *t*-test reveals that we can reject the null that the mean difference in lifespans is zero for all the sales quartiles.

Next, we compare Kaplan-Meier nonparametric survivor function estimates for farm businesses with high and low government payments as a share of sales in 1987 (the first year of the study). Figure 1 illustrates that farms in the bottom government-payments-as-a-share-of-sales quartile (*govpaycat* = 25) are less likely to survive than are those in the top quartile (*govpaycat* = 75). The Kaplan-Meier estimation does not account for the left truncation of the lifespans mentioned above, so the estimated survival probabilities are biased.⁹ However, a comparison of the survival

functions illustrates a clear difference between the groups. After 30 years, only about 53% of farms in the bottom payments-share quartile survived compared to about 64% of farms in the top quartile. Farms in the bottom payment-share quartile have an estimated mean lifespan of 30.5 years, compared to 35.9 years for farms in the top quartile. Statistical tests reveal that it is very unlikely that the survivor functions are identical across the government payment strata. Both the Savage (log-rank) test and the Wilcoxon test indicate a significant difference in the survival rates of farms that did not receive government payments in 1987 and those that did receive payments.¹⁰

A statistically significant difference between the estimated survival functions is not strong evidence that government payments influence

⁹ Survival probability estimates are biased upward as short-lived businesses are disproportionately excluded from the sample.

¹⁰ The logrank test has a chi-square statistic of 1,090.1 with an associated *p*-value less than 0.0001; the Wilcoxon test has a chi-square of 1,180.7 with an associated *p*-value less than 0.0001.

survival because other factors may be correlated with both payments and survival. For example, high-payment farms are larger on average, are more concentrated in certain types of farms and in certain regions, and are more likely to grow certain crops. If these factors are correlated with both government payments and duration of farm survival, we may observe a relationship between payments and survival that is not causal. To control for these factors we use the more general Cox proportional hazard model (Cox, 1972).

The Cox model assumes a parametric form for the effect of the explanatory variables on survival, but allows the form of the underlying survivor function to be unspecified. Cox's semiparametric model has been widely used to explain firm survival (e.g., Mata, Portugal, and Guimaraes, 1995; Audretsch and Mahmood, 1995; Disney, Haskel, and Heden, 2003). The survival time of each member of the population is assumed to follow a hazard function given by

$$(1) \quad h_{it} = h_0(t) \exp(\mathbf{x}_i' \boldsymbol{\beta})$$

where $h_0(t)$ is the baseline hazard function, x_i is a vector of explanatory variables, and $\boldsymbol{\beta}$ is a vector of parameters. To estimate $\boldsymbol{\beta}$, Cox (1972, 1975) proposed a partial likelihood function, which eliminates the unknown baseline hazard function and accounts for the fact that survival times are censored.

Table 3 presents model estimates under four alternative specifications. In all four columns, the explanatory variables include characteristics of the farm business and farm operator in the initial period, 1987. Firm characteristics include business size (logarithm of total agricultural sales), indicator variables for the SIC of the farm, the organizational structure of the farm (family-owned or otherwise), and total sales category (quartiles). In terms of operator characteristics, we use indicators for ten operator age categories, for the operator's race (white or otherwise), and the operator's main occupation (farming or otherwise).¹¹ To examine how the effect of payments varies by farm size, we interact the natural logarithm of government payments with the sales quartile indicators. The logarithm of sales is used because the distribution of government payments (like sales) is highly skewed and because

this transformation facilitates interpretation of the coefficient.¹²

In columns 2–4, we introduce 38 state fixed effects and the 24 sales–SIC interaction effects (four sales quartiles times six SIC crop categories). Column 3 introduces a control for the year in which the farm initiated production, and column 4 introduces a measure of the farm's debt-to-asset ratio. After discussing the results of these four regressions we focus on interpreting the government payments coefficient.

Across all model specifications, we find that larger enterprises are less likely to fail than smaller ones, which is consistent with studies of nonfarm businesses. We also find that hazard rates are significantly lower on farms that are family-owned, or have an operator who is male or white. The hazard rate is not significantly associated with the operator having farming as a primary occupation.

As many farm businesses fail when the operator retires, it is not surprising that being younger than seventy years old (the missing category is seventy years or older) reduces the exit hazard, and that the magnitude of this reduction in the hazard is greatest for farmers below fifty-five years. However, comparing farms operated by farmers below the normal retirement age (fifty-five years) reveals that younger farmers have a lower instantaneous probability of failure than older farmers: holding all else constant, the hazard is smallest for operators 30–34 years old and it increases gradually with age until farmers are 50–54 years old. This result does not support the hypothesis that age is positively related to financial liquidity or to the acquisition of information in a way that enhances the likelihood of survival.

In column 3 we introduce a control for the year in which the farm initiated production in order to control for policy changes that have occurred over time. Rules governing program participation and payments changed substantially over the period in which we observe farm survival (1987–97). Among other changes, planting restrictions were lifted and payments largely decoupled from production decisions with the 1996 FAIR Act. Because program rules changed, it is possible that the

¹¹ The ten age categories allow for a flexible nonlinear relationship between age and survival. This specification produced a better fit than a model using age and age-squared.

¹² The natural logarithm of government payments is set to zero when payments equal zero. As an alternative specification, we tried using government payments as a share of receipts (sales plus payments) which has the advantage of being bounded between zero and one. The main results obtained using this specification did not differ substantially from the results obtained using the logarithm of payments.

Table 3. Cox Proportional Hazard Model Estimates of Farm Business Duration under Various Specifications

Variable	(1)		(2)		(3)		(4)	
	Coeff.	SE	Coeff.	SE	Coeff.	SE	Coeff.	SE
Log Sales	-0.043	0.010	-0.061	0.010	-0.059	0.010	-0.054	0.014
SIC 111 (wheat)	-0.287	0.015	1.842	1.418	1.701	1.422	-14.620	53.368
SIC 112 (rice)	0.144	0.023	1.683	1.327	1.555	1.330	-14.427	53.365
SIC 115 (corn)	-0.157	0.012	0.838	1.229	0.963	1.233	-14.694	53.360
SIC 116 (soybean)	-0.190	0.013	1.262	0.851	1.332	0.851	-6.842	43.478
SIC 119 (cash grain)	-0.344	0.012	1.067	0.621	1.136	0.621	0.675	1.097
Operator's age <30	-0.741	0.017	-0.730	0.018	-0.850	0.018	-0.824	0.034
Operator's age 30-34	-0.777	0.017	-0.768	0.017	-0.863	0.017	-0.868	0.031
Operator's age 35-39	-0.752	0.017	-0.744	0.017	-0.840	0.017	-0.876	0.030
Operator's age 40-44	-0.716	0.016	-0.710	0.017	-0.808	0.017	-0.848	0.029
Operator's age 45-49	-0.696	0.016	-0.691	0.016	-0.791	0.016	-0.798	0.029
Operator's age 50-54	-0.626	0.015	-0.624	0.015	-0.722	0.016	-0.704	0.028
Operator's age 55-59	-0.353	0.014	-0.352	0.014	-0.435	0.014	-0.409	0.026
Operator's age 60-64	-0.115	0.013	-0.112	0.013	-0.160	0.013	-0.125	0.024
Operator's age 65-69	-0.170	0.014	-0.169	0.014	-0.203	0.014	-0.197	0.026
Sex = Male	-0.306	0.018	-0.307	0.018	-0.306	0.018	-0.294	0.033
Race = White	-0.160	0.040	-0.103	0.041	-0.104	0.041	-0.126	0.070
Organiz. = Family-owned	-0.412	0.008	-0.404	0.008	-0.409	0.008	-0.434	0.012
Main occupation = Farmer	-0.028	0.008	-0.036	0.008	-0.031	0.008	-0.078	0.016
Sales quartile ^a	Yes		Yes		Yes		Yes	
State ^a	-		Yes		Yes		Yes	
Sales quartile * SIC category ^a	-		Yes		Yes		Yes	
Initiation year ^a	-		Yes		Yes		Yes	
Debt-asset ratio	-		-		-		Yes	
Log Gov. Pay * Sales Q1	-0.033	0.002	-0.034	0.002	-0.035	0.002	0.057	0.020
Log Gov. Pay * Sales Q2	-0.048	0.002	-0.048	0.002	-0.050	0.002	-0.031	0.003
Log Gov. Pay * Sales Q3	-0.070	0.002	-0.071	0.002	-0.074	0.002	-0.046	0.004
Log Gov. Pay * Sales Q4	-0.087	0.003	-0.087	0.003	-0.090	0.003	-0.081	0.004
Log-likelihood	-1,155.888		-1,155.284		-1,154.350		-383.724	
LR chi-square (<i>p</i> -value) ^b	16,849	(<0.0001)	18,058	(<0.0001)	19,927	(<0.0001)	8,187	(<0.0001)
Maximum likelihood <i>R</i> ^{2c}	0.081		0.086		0.095		0.100	
Observations	200,187		200,187		200,187		77,594	

^aCategorical variables with more than two categories; "yes" indicates variables were included in the regression. For all categorical variables, a Wald test of the joint hypothesis that the coefficients of the categorical variables are zero is rejected at the 0.001 significance level.
^bLR chi-square is the statistic associated with the likelihood ratio test of the global null hypothesis that $\beta = 0$.
^cThe maximum likelihood *R*² is the defined $1 - \exp(-G^2/N)$ where *G*² is the likelihood ratio chi-square statistic.

effect of payments on survival was different in the 1987–95 period (call it regime 1) and in the 1996–97 period (regime 2). Similar farms have a different likelihood of spending more time in regime 2 relative to regime 1 depending on when they initiated production. In our sample, farms that initiated production at an earlier point in time would be less likely to experience regime 2 than farms that began later, *ceteris paribus*. The estimates in column 2 do not control for this policy change, so our estimates of the effect of 1987 payments on survival could be biased depending on whether the effect of payments on survival increased or decreased after 1995.

We address this potential problem by including the regressions fixed effects for the year in which the farm initiated production (there are seventy-seven different starting years). Differences between farms that were caused by the year in which the farm began farming (and hence associated with the likelihood of being in regime 1 or 2) are captured by the fixed effects. By removing these fixed effects we are effectively estimating the average effect of 1987 payments on the survival of farms. Using a Wald test we can reject the joint hypothesis that all initial-year coefficients are zero at the 0.001 significance level. However, comparing column 3 with column 2 reveals that the parameter coefficients on the variables of interest did not change much. This result is not surprising because the first regime accounts for eight of the ten years over which we observe farms surviving or exiting.

In column 4 we introduce a proxy for the farm's debt-to-asset ratio. The census collects information on the total value of land and buildings on the farm, which serves as a proxy for assets, and information on interest expenditures, which can be used to estimate farm debt.¹³ Information on the value of land and buildings and interest expenditures is only available on the "long form" of the survey, which is sent to about a third of all operations—reducing our sample to 77,594. The results in column 4 indicate that a higher debt-to-asset ratio raises the hazard rate. The coefficient associated with debt asset ratio is significant at the 5% level.

¹³ To convert interest expenditures to debt, we use the annual average thirty-year fixed rate mortgage rate (<http://www.mbaa.org/marketdata/data/02/fm30yr-rates.htm>), and assume a thirty-year loan in the fifth year of repayment. Using standard amortization methods, this implies a debt to interest expenditure ratio of 9.75 for 1987.

The coefficients associated with the logarithm of government payments interacted with sales quartiles are statistically significant and consistent across the four model specifications. To interpret these coefficients, we can rewrite (1) as

$$(2) \quad \ln h_{it} = \ln h_0(t) + \mathbf{x}_i' \boldsymbol{\beta}.$$

Let β_g be the coefficient associated with the natural logarithm of government payments ($\ln g_i$), an element of \mathbf{x}_i . It follows that

$$(3) \quad \beta_g = (dh_{it}/dg_i)(g_i/h_{it}).$$

That is, β_g is the responsiveness of the conditional probability of farm business failure to a change in government payments, expressed as an elasticity. Hence, the estimate from column 3 indicates that a 10% increase in government payments reduces the instantaneous rate of business failure by 0.35%, 0.50%, 0.74%, and 0.90%, for a representative farm in first through fourth sales quartiles, respectively.¹⁴ For example, for a farm in the fourth quartile with a hazard rate of 0.500, a 10% increase in government payments would decrease the hazard rate to 0.495.

For a more intuitive interpretation of the results, we can use the estimated parameters from the Cox partial likelihood function to estimate the survival function. For the Cox model the product-limit estimate of the survival function is

$$(4) \quad \hat{S}_i(t) = [\hat{S}_0(t)]^{\exp(\mathbf{x}_i' \hat{\boldsymbol{\beta}})}$$

where $\hat{S}_0(t)$ is the estimated baseline survivor function (see Kalbfleisch and Prentice, 1980, for details). Figure 2 displays the estimated survival functions for a representative farm with average government payments and with a 50% reduction in government payments.¹⁵ After ten years, farms receiving the average level of government payments have a chance of survival of about 35%, compared to only 25% for farms receiving half

¹⁴ In theory, farmers could respond to realized or expected government payments. Realized payments provide a noisy estimate of expected payments because a large component of realized payments is transitory. Consequently, if farmers respond to expected payments, our estimated coefficient likely underestimates the effect of a change in expected payments.

¹⁵ Government payments have fluctuated by 50% or more in consecutive years. For example, total direct payments rose from \$9.2 billion in 1992 to \$13.4 billion in 1993 and then fell to \$7.9 billion in 1994. More recently, payments fell from \$20.7 billion in 2001 to \$10.9 billion in 2002 before rising to \$17.4 billion in 2003 (<http://www.ers.usda.gov/Data/FarmIncome/finfidmu.htm>).

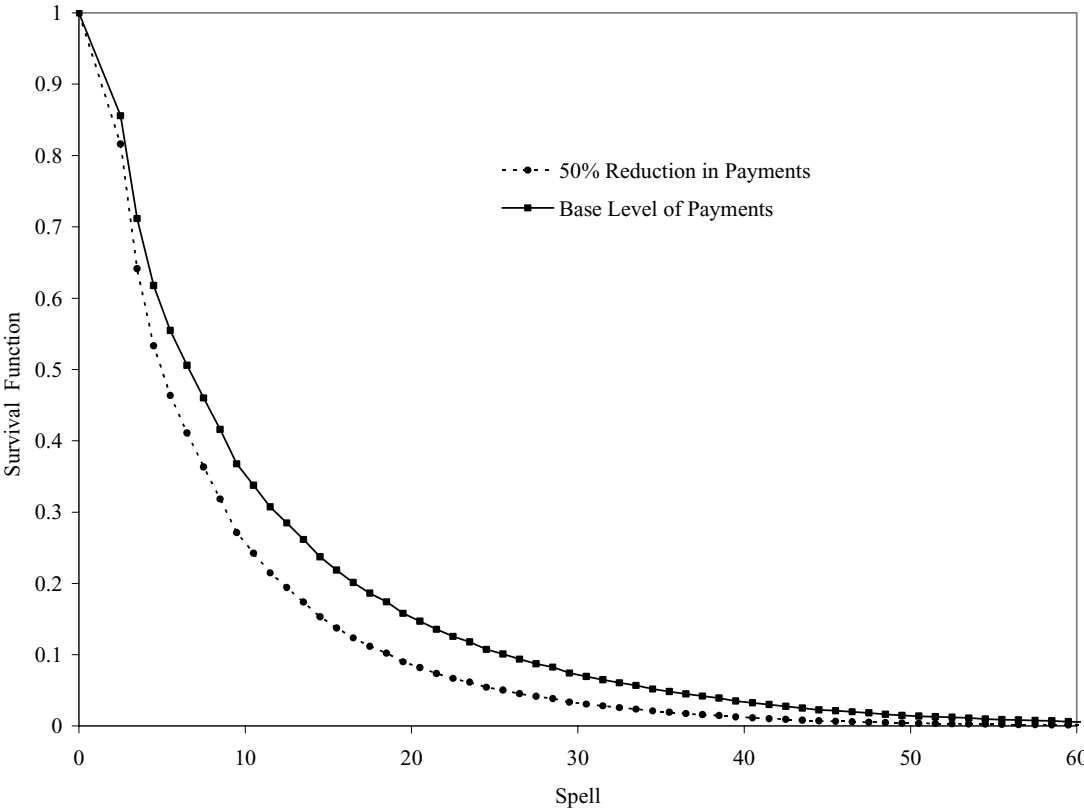


Figure 2. Product-limit survival function estimates

the average level of payments. Because the Cox model accounts for left truncation, the unbiased product-limit estimates of the survival probabilities are smaller than the biased Kaplan–Meier estimates shown in figure 1.

A large reduction in government payments could have substantially different implications for farms of different sizes. Table 4 illustrates the effect of a 50% reduction in direct government payments on expected lifespans. The effect of the payment reduction is shown separately for payment recipients and for all farms. Larger operations experience a greater reduc-

tion in life duration for two reasons. First, the marginal effect of a reduction in payments is greater for larger operations. Second, a greater percentage of large farms receive government payments (97.0% for the largest quartile, compared to 78.6% for the smallest quartile). The table shows that a 50% drop in direct government payments shortens the expected life of the largest farms by 5.4% from 14.25 to 13.48 years, and shortens the expected life of the smallest farms by 1.7% from 8.83 to 8.68 years. The positive relationship between scale and effect size is expected as farm income represents

Table 4. The Effect of a 50% Reduction in Government Payments on the Duration of Farm Businesses

Sales Quartiles	Estimated Life of Farm Business (Years)					
	Farms Receiving Payments			All Farms		
	Base	50% of Base	% Change	Base	50% of Base	% Change
Q1	9.44 (0.021)	9.24 (0.020)	−2.06	8.83 (0.020)	8.68 (0.019)	−1.71
Q2	10.93 (0.024)	10.58 (0.023)	−3.22	10.38 (0.022)	10.08 (0.022)	−2.89
Q3	12.91 (0.027)	12.32 (0.026)	−4.59	12.43 (0.026)	11.88 (0.025)	−4.38
Q4	14.67 (0.031)	13.86 (0.029)	−5.53	14.25 (0.030)	13.48 (0.029)	−5.41

a larger share of total farm household income for larger farms (Mishra et al., 2002).

Conclusions

Government payments have a small but statistically significant positive effect on farm business survival. This finding could be explained by several factors. Farms receiving relatively high payments may be able to bid up the price of land and other fixed resources—causing farms receiving lower payments to exit. Government payments may also relieve liquidity constraints allowing farms receiving more payments to achieve a more efficient scale and remain in business longer. Additionally, higher payments may make farming more profitable relative to alternative occupations, thereby reducing incentives to exit agriculture.

The study also found that government payments increase business survival rates proportionally more for larger farms. This result is probably attributable to the fact that government payments' share of farm household income increases with total sales. While payments appear to disproportionately benefit larger operations, the long run consequences of an increase in payments for agricultural structure are ambiguous. Because the study did not account for the size of farms entering production, it is not possible to conclude that lower failure rates for larger farms necessarily increase the concentration of production. Further work would be needed to understand how government payments influence the size distribution of farm businesses.

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References

- Atwood, J., M. Watts, and A. Baquet. 1996. "An Examination of the Effects of Price Supports and Federal Crop Insurance upon the Economic Growth, Capital Structure, and Financial Survival of Wheat Growers in the Northern High Plains." *American Journal of Agricultural Economics* 78(February):212–24.
- Audretsch, D. 1991. "New-Firm Survival and the Technological Regime." *The Review of Economics and Statistics* 73(3):441–50.
- Audretsch, D., and T. Mahmood. 1995. "New-Firm Survival: New Results Using a Hazard Function." *The Review of Economics and Statistics* 77(1):97–103.
- Baldwin, J., and P. Gorecki. 1991. "Firm Entry and Exit in the Canadian Manufacturing Industries." *Canadian Journal of Economics* 24(2):300–23.
- Bates, T. 1990. "Entrepreneur Human Capital Inputs and Small Business Longevity." *The Review of Economics and Statistics* 72(4):551–59.
- Becker, E. 2001. "Far from Dead, Subsidies Fuel Big Farms." *New York Times*, Section A, p. 1, May 14.
- Cox, D.R. 1972. "Regression Models and Life-Tables (with Discussion)." *Journal of the Royal Statistical Society, Series B* 34:187–220.
- . 1975. "Partial Likelihood." *Biometrika* 62(2):269–76.
- Disney, R., J. Haskel, and Y. Heden. 2003. "Entry, Exit, and Establishment Survival in UK Manufacturing." *Journal of Industrial Economics* 51(1):91–112.
- Dixon, B.L., N. Ma, B.L. Ahrendsen, L. Settlege, and J. Stam. 2004. "Factors Affecting State-Level Chapter 12 Filing Rates: A Panel Data Model." *Emory Bankruptcy Developments Journal* 20(2):401–26.
- Dunne, T., M. Roberts, and L. Samuelson. 1988. "Pattern of Entry and Exit in the US Manufacturing Industries." *RAND Journal of Economics* 19(4):495–515.
- Ericson, R., and A. Pakes. 1992. "Markov-Perfect Industry Dynamics: A Framework for Empirical Work." *Review of Economic Studies* 62(1):53–82.
- Evans, D. 1987a. "The Relationship between Firm Growth, Size, and Age: Estimates for 100 Manufacturing Industries." *Journal of Industrial Economics* 35(June):567–81.
- . 1987b. "Tests of Alternative Theories of Firm Growth." *Journal of Political Economy* 95(Aug.):657–74.
- Evans, D., and B. Jovanovic. 1989. "An Estimated Model of Entrepreneurial Choice under Liquidity Constraints." *Journal of Political Economy* 97(4):808–27.
- Hallam, A. 1993. "Empirical Studies of Size, Structure, and Efficiency in Agriculture." In A. Hallam, ed. *Size, Structure, and the Changing Face of American Agriculture*. Boulder, CO: Westview Press, pp. 204–31.
- Holtz-Eakin, D., D. Joulfaian, and H. Rosen. 1994. "Sticking It Out: Entrepreneurial Survival and Liquidity Constraints." *Journal of Political Economy* 102(1):53–75.
- Hubbard, G. R. 1998. "Capital-Market Imperfections and Investment." *Journal of Economic Literature* 36:193–225.
- Huffman, W.E., and R.E. Evenson. 2001. "Structural and Productivity Change in

- US Agriculture 1950–1982.” *Agricultural Economics* 24(2):127–47.
- Jovanovic, B. 1982. “Selection and Evolution of Industry.” *Econometrica* 50(May):649–70.
- Kalbfleisch, J.D., and R.L. Prentice. 1980. *The Statistical Analysis of Failure Time Data*. New York: John Wiley and Sons, Inc.
- Key, N., and M.J. Roberts. 2005. “Financial Market Imperfections and Structural Change in Agriculture.” In Ami R. Bellows, ed. *Focus on Agricultural Economics*. New York: Nova Science Publishers, Inc., pp. 29–51.
- Kimhi, A., and R. Bollman. 1999. “Family Farm Dynamics in Canada and Israel: The Case of Farm Exits.” *Agricultural Economics* 21(1):69–79.
- Leathers, H. 1992. “The Market for Land and the Impact of Farm Programs on Farm Numbers.” *American Journal of Agricultural Economics* 74:291–98.
- Mata, J., P. Portugal, and P. Guimaraes. 1995. “The Survival of New Plants: Start-Up Conditions and Post-entry Evolution.” *International Journal of Industrial Organization* 13:459–81.
- Mishra, A., H. El-Ostra, M. Morehart, J. Johnson, and J. Hopkins. 2002. “Income, Wealth, and the Economic Well-Being of Farm Households.” Agricultural Economic Report No. 812, U.S. Department of Agriculture, ERS, Washington DC, July.
- Nelson, E.B. 2002. “Payment Limitations Amendment Passes Senate: Nelson Co-sponsored Amendment Restores Equity to Farm Payments.” Press release, February 7. Available at <http://bennelson.senate.gov/news/details.cfm?id=240851&&>
- Pakes, A., and R. Ericson. 1998. “Empirical Implications of Alternative Models of Firm Dynamics.” *Journal of Economic Theory* 79(1): 1–46.
- Shepard, L., and R. Collins. 1982. “Why Do Farmers Fail? Farm Bankruptcies 1910–78.” *American Journal of Agricultural Economics* 64:609–15.
- Sumner, D.A., and J.D. Leiby. 1987. “An Econometric Analysis of the Effects of Human Capital on Size and Growth among Dairy Farms.” *American Journal of Agricultural Economics* 69(2):465–70.
- Taylor, M.P. 1999. “Survival of the Fittest? An Analysis of Self-employment Duration in Britain.” *Economic Journal* 109(454):140–55.
- Tweeten, L. 1993. “Government Commodity Program Impacts on Farm Numbers.” In A. Hallam, ed. *Size, Structure and the Changing Face of American Agriculture*. Boulder, CO: Westview Press, pp. 336–64.
- U.S. Department of Agriculture (USDA-ERS). Historical Commodity Specific Yearbooks (various years) Economic Research Service. www.ers.usda.gov
- Weiss, C. 1999. “Farm Growth and Survival: Econometric Evidence for Individual Farms in Upper Austria.” *American Journal of Agricultural Economics* 81(February):103–16.
- Williams-Derry, C., and K. Cook. 2000. “Green Acres: How Taxpayers Are Subsidizing the Demise of the Family Farm.” Environmental Working Group, Washington DC, April.
- Zepeda, L. 1995. “Asymmetry and Nonstationarity in the Farm Size Distribution of Wisconsin Milk Producers: An Aggregate Analysis.” *American Journal of Agricultural Economics* 77(4):837–52.